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Welfare Reform and the Labor Market:
Earnings Potential and Welfare Benefits in California, 1972–1994

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Abstract

Promotion of work is prominent in the rhetoric of current welfare reform efforts. The success of welfare-to-work policies is in part dependent on earnings available in employment. In this paper we use Current Population Survey data for the years 1972–1994 to develop measures of potential earnings from full-time work for low-skilled men and women in California and to compare the trend in earnings capacity for such people to welfare benefits. We find that while benefits have declined, earnings capacity has fallen faster, and the downward trend is particularly pronounced for men. Both the downward trends in benefits and potential earnings appear to have accelerated in recent years. State attempts to address the problem of low wages by expanding the opportunity for combining welfare with work may conflict with federal efforts to require that assistance be transitory.

Welfare Reform and the Labor Market:
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Promotion of work is prominent in the rhetoric of current welfare reform efforts. One of the stated purposes of the block grant system for Temporary Assistance for Needy Families (TANF) established by the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 (PRWORA) is to "end the dependence of needy parents on government benefits by promoting job preparation, work, and marriage."[1] In a similar spirit, California's *Proposed Redesign of the Welfare System* calls for "changing the overall purpose of welfare . . . to provide time-limited assistance to families in crisis" and "achieving independence through work while at the same time strengthening and supporting children and families" (California Department of Social Services 1996: 2).

Talking about work in connection with welfare is not new. The novelty in the current debate is the emphasis placed on devolution of responsibility for design and conduct of work-oriented policy to the states and upon making assistance short-term. Both federal and state policy makers seem to believe that with relatively modest effort households in need can be moved to self-support. This is to be accomplished in part by making use of public assistance less attractive than work. Again in the words of California policy makers, "the benchmark against which welfare eligibility and benefit packages will be structured is fairness and equity with low-income working families. This is consistent with the principle that welfare should not pay more than work" (California Department of Social Services 1996: 2). We term this California principle the *equity benchmark*.[2]

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[1]Conference report on H.R. 3734, Personal Responsibility and Work Opportunity Reconciliation Act of 1966, Section 401(a)(2). *Congressional Record*, July 30, 1996, H8831.

[2]The state's plan does not include a sufficiency benchmark. The equity principle is a restatement of the "least benefit" principle of the English Poor Law of 1834. The fact that earnings from employment exceed the welfare stipend does not guarantee that work will be more attractive, since time and other costs are presumably also involved in the labor-supply calculus.

Success in attaining the goals of making assistance temporary and satisfying the equity benchmark is very much dependent upon circumstances in the labor market. This paper develops measures of potential earnings or earnings capacity of women and men who work full time in California, and we compare earnings trends to trends in welfare benefits. We find that earnings capacity has declined for men and women with characteristics associated with risk of need for public assistance and that this decline has been particularly sharp since the late 1980s. We argue that recent welfare policies adopted in California to encourage recipients to combine work with welfare make sense in the context of what seems to be happening to earnings, but that desirable policy from California's perspective may be inconsistent with time limits imposed by PRWORA.

TRENDS IN WELFARE AND WAGES

We begin by looking at data on welfare benefits and the earnings of women and men.

For our purpose, welfare benefits include the AFDC payment and the value of the Food Stamps that come with AFDC. We ignore for the time being benefits received from other programs, although we will return to them at the end of the paper.[3] The data are plotted in Figure 1.

The message of Figure 1 is that between 1972 and 1990 real benefits changed very little. For a family of three, the value of AFDC and Food Stamp benefits averaged about $11,000 per year in 1994 dollars. Since 1990, the state has cut back benefits substantially, so that the average (July) benefit level for 1993–95 is 13 percent lower than the average over the period 1988–90. Thus if earnings potentially

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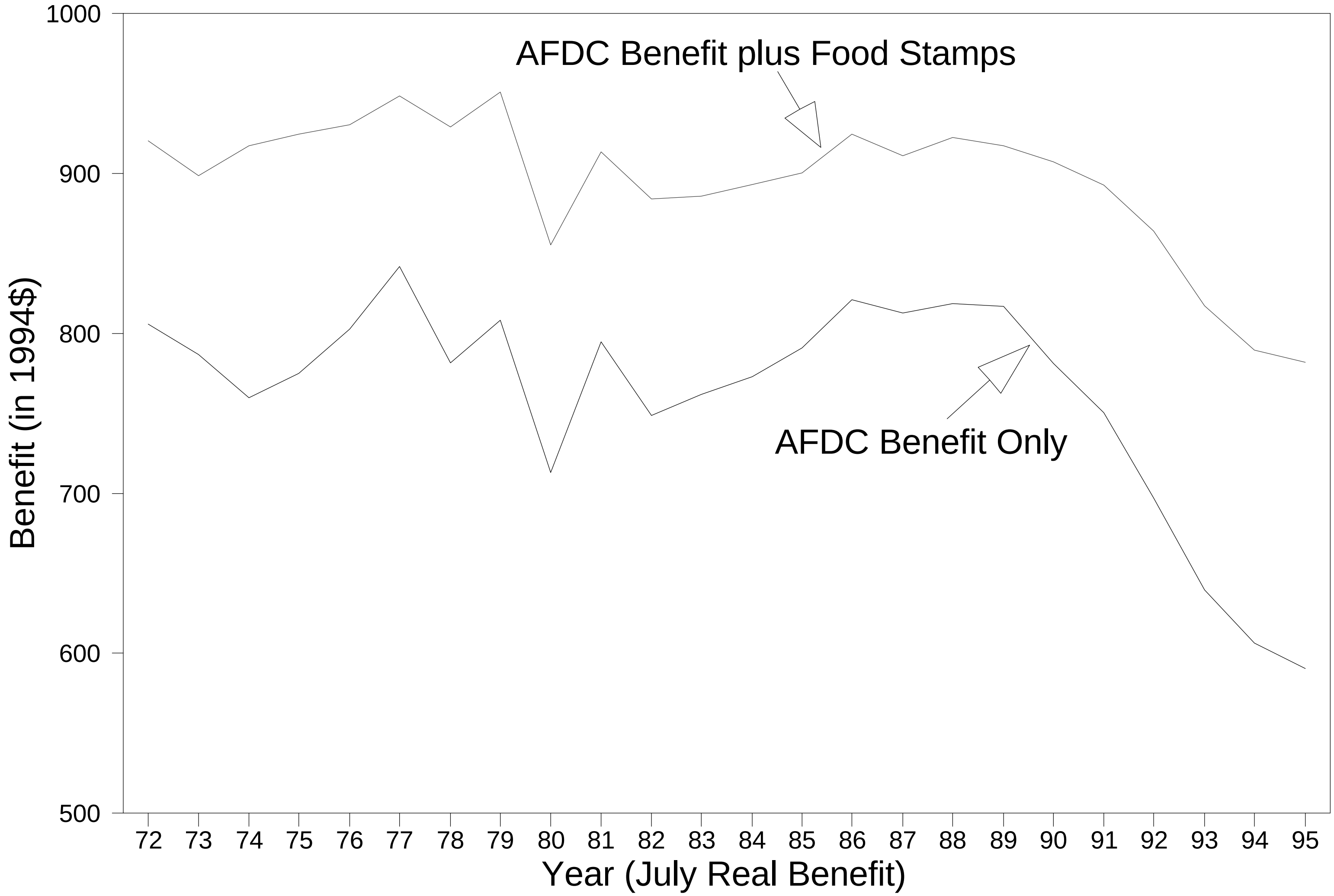
[3]All AFDC recipients are categorically eligible for Medi-Cal. Medi-Cal is California's version of Medicaid, the medical insurance program provided for AFDC recipients and certain other low-income families. Because Medi-Cal covers virtually all common health costs, the AFDC and Food Stamps benefit used in this analysis is deflated by the consumer price index for expenditures other than medical care. Food Stamp benefits are calculated assuming the maximum allowable shelter allowance and no other household income.

FIGURE 1

California's Welfare Benefit, 1972-1995

Family of three without other income



available from work on the outside have not declined or improved, the benefit trend has raised the likelihood that the state's system meets the equity standard, that is, that welfare pays less than work.

Data on welfare benefits are available from administrative sources. Data on potential earnings are more difficult to come by. In this paper we base our analysis of trends in potential earnings on income from work reported by California women and men in the March Current Population Survey (CPS) for each year from 1973 to 1995. The March surveys include questions about earnings and other sources of income for individuals in the preceding year as well as total weeks worked and hours typically worked per week. The sample is sizable and approximately representative (when appropriately weighted) of the state's population. For our purposes, a major shortcoming of the data is the absence of precise information on the hourly wage. The usual approach is to construct an hourly wage by dividing reported annual earnings by the product of hours typically worked per week and total weeks worked. Since all three variables are measured with error, the result is of questionable utility.

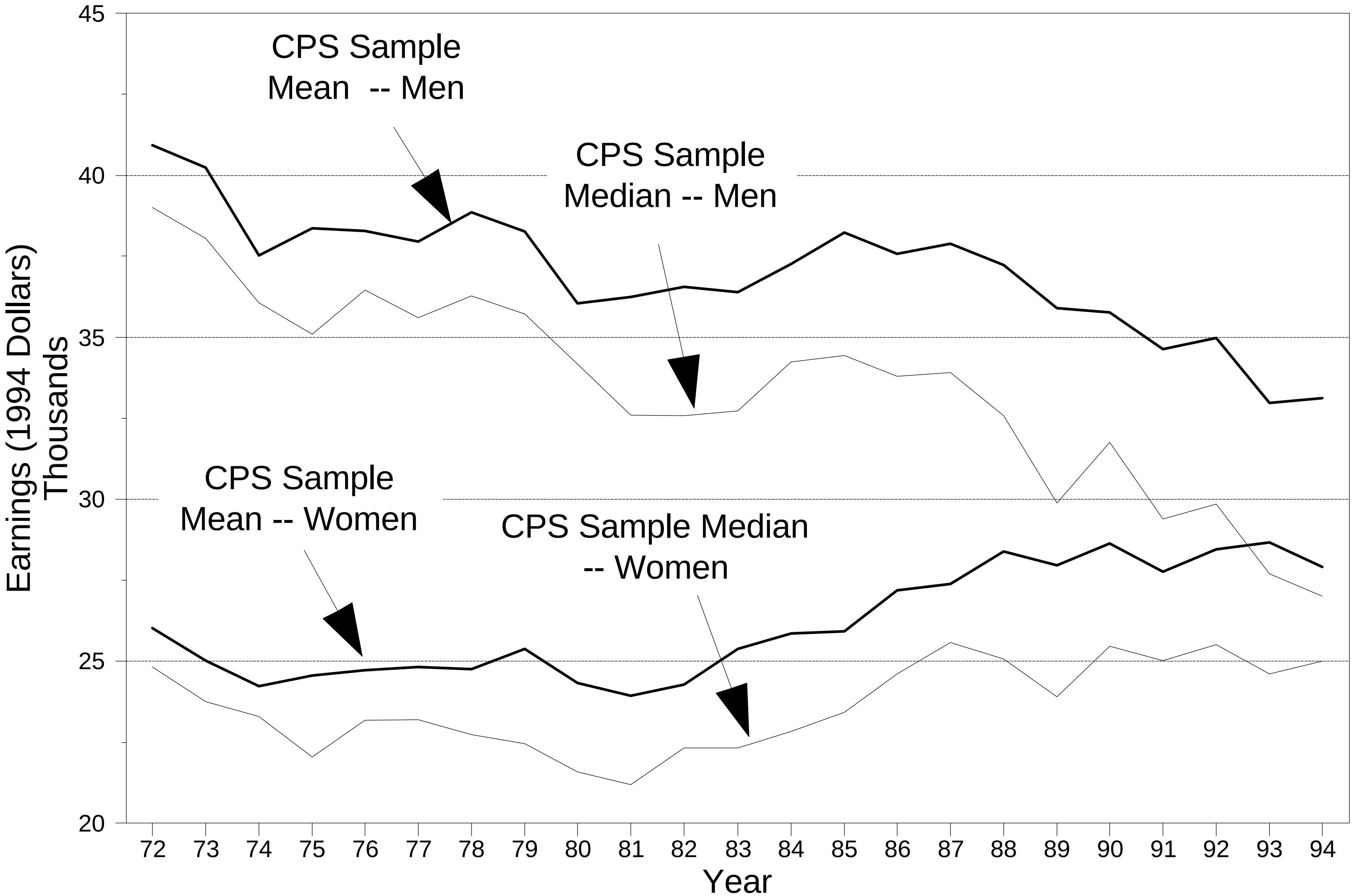
To avoid errors engendered by the absence of hourly wage data, we take as a first indicator of the trend in earnings potential the reported earnings of all civilian men and women age 18–44 who claim to have worked full time for the entire year preceding the survey. We define "full time" to mean working 35 or more hours per week; "all year" means 50 or more weeks (including paid vacations). Since the earnings distribution tends to be skewed, we calculate both the (weighted) mean and median earnings for our sample from each survey. Given the retrospective character of the survey, this means we have earnings for calendar years 1972 through 1994. The results appear in Figure 2.

At least four features of these plots seem important. First, for both men and women it appears that the degree of skewness in the dispersion of earnings, as reflected in the difference between the mean and median of the sample distribution, is increasing. Second, at least since the early 1980s, the earnings of women who work full time have been rising relative to the earnings of their male counterparts. Within the CPS the 1977–79 median for full-time women workers averaged 64 percent of the median for men;

FIGURE 2

Earnings of Men and Women Working Full Time

California,1972-1994



by 1992–94 this percentage had risen to 89 percent. Third, the gain in relative earnings has been accomplished both because women's earnings have been rising and because the real earnings of men have been falling. Finally, the medians are substantial—approximately $28,000 for men and $25,000 for women in 1994.

At first pass and viewed from the perspective of the welfare reform objectives cited earlier, there is good and bad news here. On the good side, if all women are experiencing the upward trend in earnings evident in the figure, potential earnings would appear to be rising relative to welfare benefits. On the bad side, if the potential earnings of fathers of children at risk of poverty follow the trend evident for all men, the attractiveness to women of marriage as a means of gaining economic resources is diminishing.

However provocative, caution must be exercised in drawing conclusions about labor market trends from these simple plots. The raw earnings data from the CPS may be very misleading as indicators. There are two culprits in this deception: mixture and selection.[4] Mixture involves the changing composition of the workforce. Over the 23 years covered by our CPS data, California's population has changed substantially in ways certain to affect potential earnings. In particular, the distribution of the population by age, education, race, and location has altered. If levels and trends in potential earnings differ across subgroups defined by such variables, then year-to-year changes in average earnings may reflect a changing population mix and not variation in the prospects for individuals.

The selection problem is obvious: Since 1972 labor force participation has changed dramatically for women and substantially for men. In 1972–74 about 26 percent of women aged 18–44 in the state reported working full time, full year by our definition. By 1992–94 this had increased to 42 percent. The corresponding change for men is from 57 to 65 percent. Many of the remainder worked, but not full time. Since it is likely to be good wages that attract and enable workers to stay in the same job all year,

[4]We should include precautions concerning the incidence of nonreporting and census procedures for imputation of missing earnings data as well, since both have changed over time. Education was not included among the variables used to impute missing income values until 1976 (Lillard et al. 1986).

inferences drawn from trends in earnings only for full-time, full-year workers are undoubtedly biased as an indicator of trends in potential earnings for the entire population. This problem is exacerbated by cyclical variation in the ability of workers to avoid joblessness.

One additional problem pertinent to our concern with the relationship between trends in earnings and trends in welfare should be noted. Welfare payments are not subject to income taxation or withholding for unemployment insurance and social security. To the extent that tax liability has varied over the period we discuss, actual trends in take-home pay may differ from those evident in Figure 2.

ESTIMATING POTENTIAL EARNINGS

To examine trends in earnings for workers holding characteristics constant, we have developed a measure of *gross earnings capacity* (GEC) by year by regressing earnings for each full-time, full-year worker in our data on a set of variables reflecting location and personal characteristics believed to affect what people are paid for working. To correct for selection bias, we use the two-stage correction procedure first proposed by Heckman (1976).[5] Comparison of selection-adjusted predicted earnings from year to year given the same characteristics permits isolation of the influence of general wage trends. To recognize the consequences for take-home pay brought about by changes in tax law, a measure of *net earnings capacity* (NEC) is developed by subtracting net tax liability (as reduced by the Earned Income Tax Credit) from GEC. The procedure is repeated for each year's data. In this section we review the selection adjustment procedure, estimation of the earnings function, and the adjustment made to obtain the NEC.

Since the Heckman adjustment procedure is well-known, we provide only a brief summary that emphasizes application in this context. In the first stage of the procedure, the probability of being a full-time worker (FTYR) is estimated using a probit model and data for all noninstitutionalized civilian

[5]This method of selection adjustment relies on very stringent functional form assumptions. For discussions, see for example Goldberger (1983) and Stolzenberg and Relles (1990).

California residents meeting our age restriction. Independent variables employed in this model include location of residence (central city, remainder of Metropolitan Statistical Area, outside of MSA), race, education, age, health status, marital status, and number of children. From this equation, λ (termed the inverse Mills ratio) is estimated for each individual, where

$$\lambda = \frac{\phi(\gamma' w)}{\Phi(\gamma' w)} \tag{1}$$

In equation (1), ϕ and Φ are, respectively, the density and cumulative distribution function of a standard normal variable, w is a vector of independent variables, and γ is a vector of estimated coefficients.

In the second stage, the sample is restricted to full-time workers, and earnings are estimated using ordinary least squares (OLS). The estimated earnings equation is of the form

$$LOGEARN = X\beta + c\lambda + \varepsilon \tag{2}$$

where LOGEARN is the logarithm of observed earnings, X includes demographic characteristics that may affect earnings, λ is the Heckman selection correction term (see equation (1)), and ε is the random error term distributed $N(0,\sigma^2)$. Independent variables included in X are location of residence, race, education, and age. The inclusion of λ compensates for the increased conditional mean earnings in the truncated sample and thus yields consistent estimates of β.

Given the regressions for each year 1972–94, earnings capacity was calculated in two steps. In the first step, a vector of demographic characteristics for a representative individual, $\overline{X}$, was selected. These characteristics and the parameter estimates were then combined to produce expected logarithm of earnings for this representative case. In the second step, expected logarithm of earnings was transformed into dollar amounts. Given the assumption we have made for the distribution of ϵ in equation (3), taking the antilog of the predicted logarithm of earnings gives predicted *median* gross earnings.[6] To compare our results to those from studies using *mean* earnings, we also compute mean predicted gross earnings by multiplying median predicted gross earnings by $\exp(\frac{1}{2}\hat{\sigma}^2)$, where $\hat{\sigma}$ is the estimated standard error of the regression.[7] Nominal earnings are converted to 1994 dollars using the CPI-U. To summarize:

$$Median\ \ G\hat{E}C_t \ = \ \exp(\bar{X}\hat{\beta}_t) \ \times \ \frac{CPI_{1994}}{CPI_t} \tag{3}$$

$$Mean\ \ G\hat{E}C_t \ = \ [\exp(\bar{X}\hat{\beta}_t) \ \times \ \exp(\frac{1}{2}\hat{\sigma}_t^{\,2})] \ \times \ \frac{CPI_{1994}}{CPI_t} \tag{4}$$

Since taxes may affect returns to working, we also calculate net earnings. Net earnings capacity (NEC) is defined as GEC minus estimated net taxes due at that level of earnings for a three-person family with no other income. Net taxes are estimated federal income tax (FEDTAX) and Social Security tax (FICA) liability minus the Earned Income Tax Credit (EITC). We ignore California state income tax liability because at the levels of earnings considered here a single-parent family has no state income tax liability. To summarize:

$$Median\ \ N\hat{E}C \ = \ [\exp(\bar{X}\hat{\beta}) \ - \ FEDTAX \ - \ FICA \ + \ EITC] \ \times \ \frac{CPI_{1994}}{CPI_t} \tag{5}$$

$$Mean\ \ N\hat{E}C = [[\exp(\bar{X}\hat{\beta}) \times \ \exp(\frac{1}{2}\hat{\sigma}^2)] - FEDTAX - FICA + EITC] \ \times \ \frac{CPI_{1994}}{CPI_t} \tag{6}$$

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[6]The fact that the antilog of predicted logarithm of earnings is the conditional median, and not the conditional mean, expected value of earnings is still occasionally ignored. In evaluating trends, as we want to do, the difference between the conditional median value of earnings and the conditional mean value acquires significance if there is a time trend in the estimated value of σ^2.

[7]This method for getting the conditional mean predicted value is presented in Goldberger (1968).

Note that the estimate of mean net earnings capacity reflected in (6) is biased because $\hat{\sigma}^2$ is not estimated from a sample of observations on net earnings. If on balance the tax system is progressive, the bias will be to exaggerate the difference between predicted median and mean NEC.

RESULTS FOR WOMEN

We present our results for women in detail. Procedures for the estimating earnings capacity for men are identical.

The Earnings Regression

We use the results from the 1995 CPS to illustrate our procedure. The data are from the 1995 CPS and therefore cover 1994. Appendix 1 provides a description of the variables in the 1995 CPS probit and earnings regressions. This list is representative of the variables used over the entire time period, although some definitions were not uniform over all 23 years due to changes in the CPS survey. The availability of the data is noted in the definition.

Tables 1 and 2 provide a sample of the results of the two-step estimation procedure. Table 1 shows the results from the first-stage probit estimation, and Table 2 shows the results of the second-stage OLS estimation. Complete results for all 23 years are available from the authors upon request.

We next calculate expected earnings for a representative woman. By “representative” we mean that variables for which effects are not explicitly displayed are set at levels judged to reflect the circumstances of women age 15–44 at the time of the survey and at risk of receiving welfare. The

TABLE 1

**Female Probit Equation Estimation Results:
Probability of Full-Time, Full-Year Employment, 1995**

Variable	Mean	Std. Dev	Coefficient	Std. Error
CITY	0.3579	0.4795	0.2787	0.1751
SUBURB	0.5693	0.4953	0.3169*	0.1723
NOTID	0.0331	0.1789	0.1337	0.2440
BLACK	0.0712	0.2572	0.1247	0.1319
ASIAN	0.0878	0.2830	0.0827	0.0989
HISPANIC	0.3504	0.4772	0.2120	0.0719
OTHER	0.0111	0.1046	-0.0152	0.2534
ED	12.2215	3.4776	0.0472***	0.0172
HSGRAD	0.7799	0.4144	0.3283***	0.1176
COLLGRAD	0.2017	0.4013	0.0920	0.0988
AGE	32.4508	7.1732	0.1325***	0.0322
AGE24	8.8600	6.5399	-0.0754**	0.0381
AGE34	2.3065	3.2546	-0.0640***	0.0195
MARRIED	0.5738	0.4946	-0.2166***	0.0758
PREVMAR	0.1458	0.3530	0.0074***	0.0995
NUMCHILD	1.278	1.300	-0.2467***	0.0265
HEALTHPG	0.0220	0.1466	-0.3967**	0.1935
CONSTANT	1.0000	0.0000	-4.4945***	0.7700

Dependent Variable: FTYR
Number of Observations: 2314
Mean of Dependent Variable: 0.4239
Standard Deviation of Dependent Variable: 0.4943

*** Significant at the 1% level; ** significant at the 5% level; * significant at the 10% level.

TABLE 2

Female Earnings Regressions Results:
Earnings from Full-Time, Full-Year Work, 1995
(Heckman Selection Model)

Variable	Mean	Std. Dev	Coefficient	Std. Error
CITY	0.3512	0.4776	0.1690	0.1173
SUBURB	0.5941	0.4913	0.2139*	0.1159
NOTID	0.0249	0.1559	-0.1007	0.1613
BLACK	0.0869	0.2819	-0.1267	0.0709
ASIAN	0.0991	0.2989	-0.0815	0.0557
HISPANIC	0.2779	0.4482	-0.1664***	0.0410
OTHER	0.0120	0.1089	-0.1507	0.1419
ED	13.2892	2.7290	0.0314**	0.0153
HSGRAD	0.8942	0.3077	0.4417***	0.0866
HSINT	1.7365	1.7819	0.1028***	0.0314
COLLGRAD	0.2829	0.4506	0.0104	0.0996
COLLINT	0.0942	0.3431	-0.0589	0.0655
EXPER	14.4610	6.9836	0.0450***	0.0090
EXP2	257.8383	206.6696	-0.0007**	0.0003
CONSTANT	1.000	0.0000	8.4263***	0.2317
LAMBDA	—	—	0.0903	0.0630

Dependent Variable: LGEARN
Number of Observations: 943
Mean of Dependent Variable: 10.0631
Std. Deviation of Dependent Variable: 0.6140
Log Likelihood: -2048.03
R^2 (unadjusted): .33

*** Significant at the 1% level; ** significant at the 5% level; * significant at the 10% level.

variable values used in this calculation are summarized in Table 3. We assume the representative woman reports never being married, since this is the circumstance of most women receiving AFDC.

Data from the 1980 census were used to calculate variable values for residential and race/ethnic classification. The variable values are an estimate of the proportion of females at risk that are in each classification. The number of women at risk by classification was estimated by multiplying the proportion of each classification with income under 125 percent of poverty by the number of females in each classification age 15–44. Residential classifications used in the regression are urban (individuals living in a central city), suburban (individuals living in the balance of the MSA), and rural (individuals living in an area outside an MSA). The estimates are 0.35 for urban, 0.60 for suburban, and 0.05 for rural. Race/ethnic classifications used in the regression are white, black, Hispanic, Asian, and other. The estimates are 0.45 for white, 0.15 for black, 0.34 for Hispanic, 0.06 for Asian, and 0 for other (less than 2 percent of the sample).

The census offers no tabulation of persons by education and poverty status, so we used the finding from Burtless (1994) that approximately half the 25-year-old women who reported receiving welfare in the year prior to each survey from the National Longitudinal Study of Youth for the period 1979–1990 reported having competed high school and none had a four-year college degree. Therefore, the earnings predictions we use are an average of log predicted wages of an individual with 11 years of education and an individual with a high school degree. In all cases the representative individual is assumed to be 25 years old.

Figure 3 presents the results of calculation of mean and median GEC from the regression results for each year’s CPS sample (see equations (3) and (4)) for the representative woman. In Figure 4 the CPS sample median (see Figure 2) is plotted with the predicted median for the representative woman.

Once again, there are things to note.

TABLE 3

Variable Weights Used in Calculating Probability of Full-Time Employment and Representative Gross Earning Capacity

Variable	Weight in Probability Equation	Weight in Earnings Estimate
INTERCEPT	1	1
CITY	0.35	0.35
SUBURB	0.60	0.60
NOTID	0	0
BLACK	0.16	0.16
ASIAN	0.05	0.05
HISPANIC	0.34	0.34
OTHER	0	0
ED	11.5	11.5
HSGRAD	0.5	0.5
HSINT	NA	0
COLLGRAD	0	0
COLLINT	NA	0
AGE	25	NA
AGE24	1	NA
AGE34	0	NA
EXPER	NA	7.5
EXP2	NA	56.25
MARRIED	0	NA
PREVMAR	0	NA
NUMCHILD	2	NA
HEALTHPG	0	NA

FIGURE 3

Mean and Median Gross Earning Capacity

Representative California Woman, 1972-1994

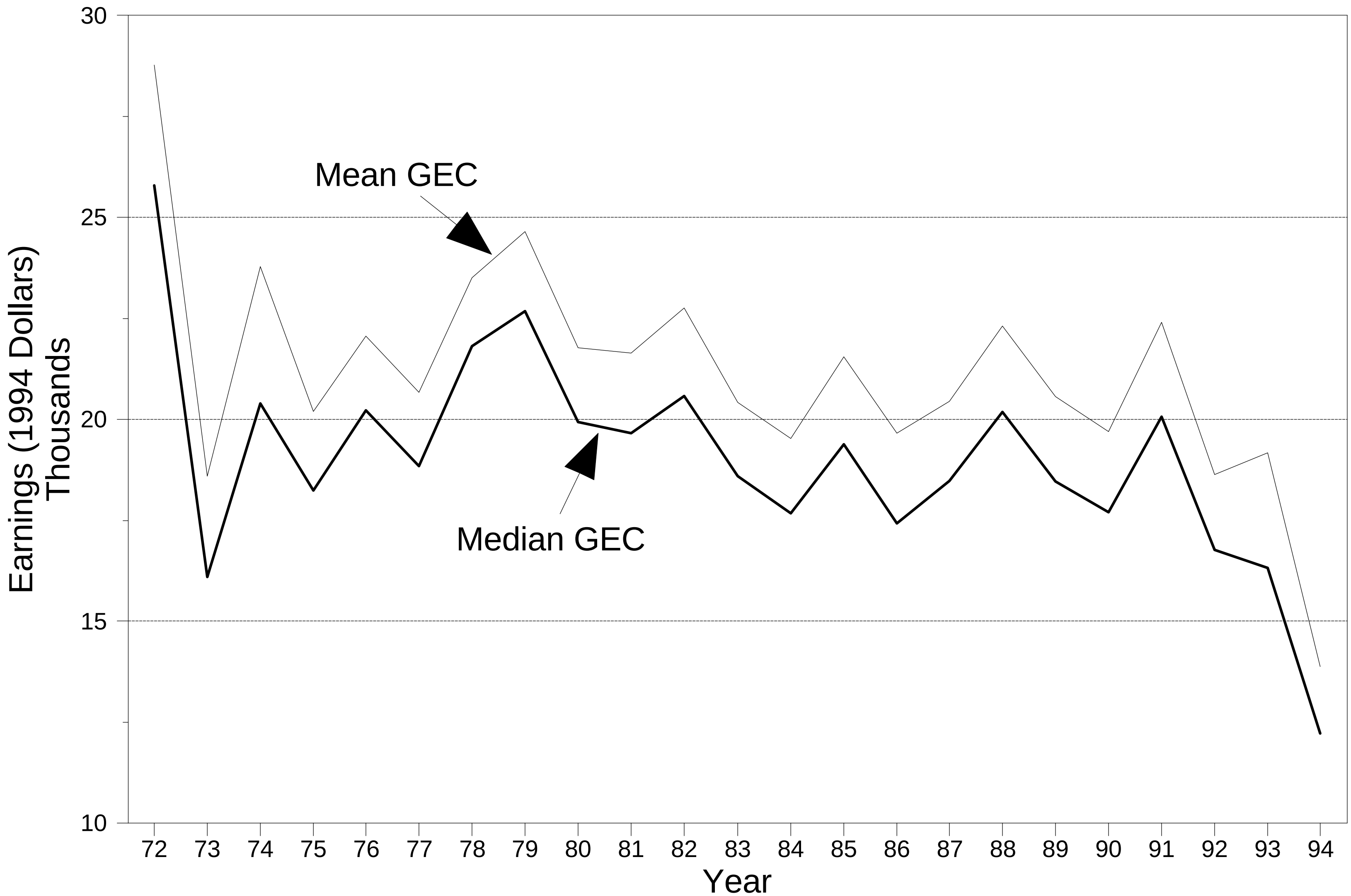
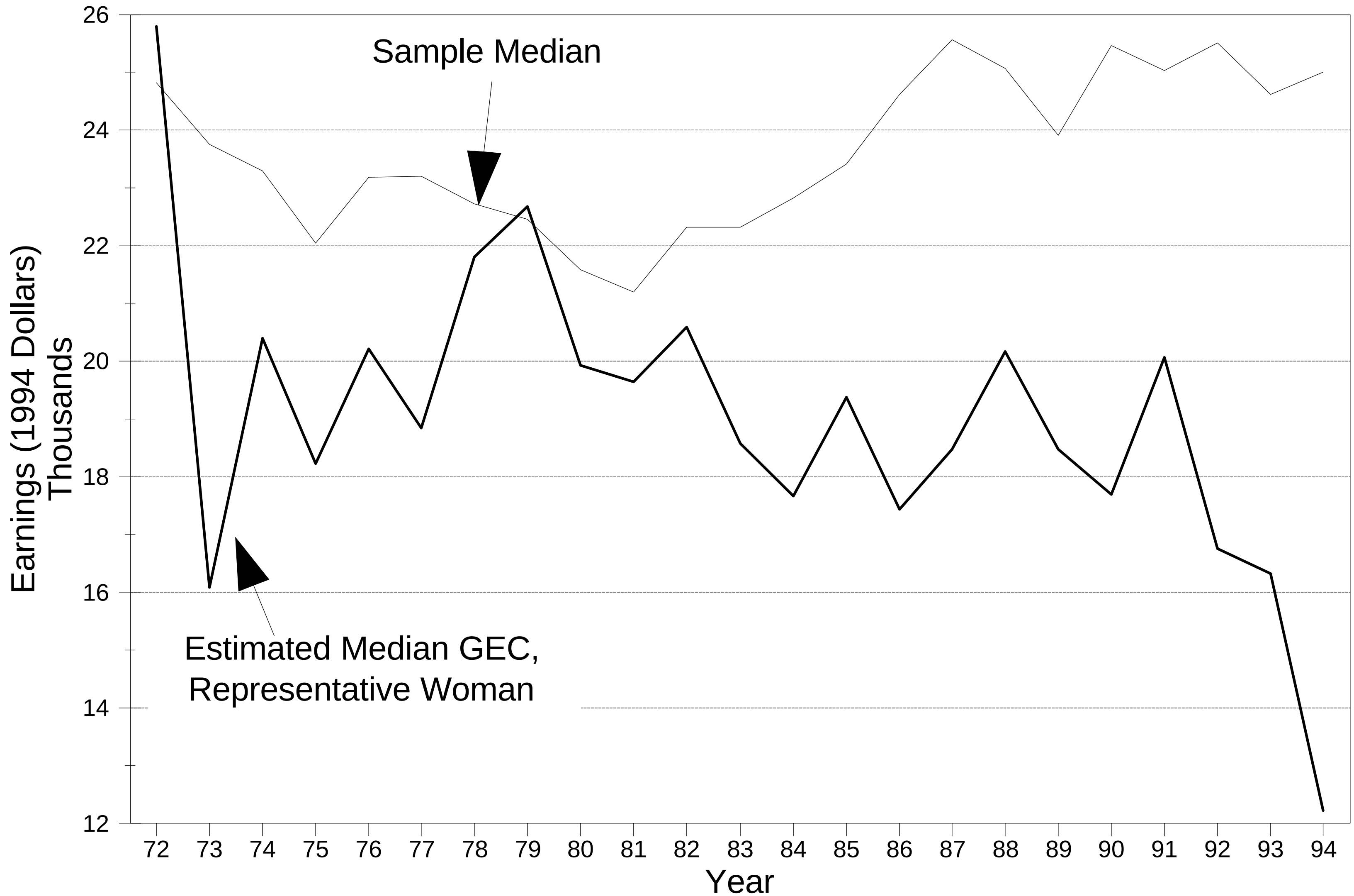


FIGURE 4

Estimated Median GEC and Median Earnings

California Women Working Full Time, 1972-1994



First, the GEC estimates for 1972 and 1994 are anomalous, with the 1972 number implausibly high and the change between 1993 and 1994 implausibly great. We have conducted extensive analyses of both estimates, and we can find no error. A contributing factor to the high 1972 estimate appears to be exceptional earnings among persons in the March 1973 sample with less than a high school education. In samples subsequent to 1973, reported earnings are much lower among persons reporting this level of educational attainment. The decline in 1994 seems to be supported by examination of raw means and medians by subgroups: all median earnings by education go down except for those with college degrees.

Second, the data for average earnings of full-time, full-year workers and the estimates for GEC present a substantially different picture of the California labor market when viewed from the perspective of women likely to be at risk of need for public assistance. Calculation of trends is sensitive to choice of endpoints; here we use 1975 and 1993 as two points similarly situated in the business cycle that avoid the possible anomaly in 1972–73 and 1994. Between 1975 and 1993 median earnings for women in the CPS sample working full time increased by 11.7 percent, while estimated median GEC for our representative woman fell by 10.5 percent and real benefits declined 11.6 percent. Thus, over the long run, the decline in benefits has exceeded the decline in earnings capacity. However, recent experience has been much different. Between 1987–89 and 1992–94, real benefits fell by 10 percent, while estimated earnings capacity declined by 19 percent.

Third, unlike the results for the sample as a whole, the difference between predicted mean and median GEC is relatively constant over time, which is another way of saying that the standard errors of the selection-adjusted earnings estimates are relatively stable (see equation (4)). There is a modest negative correlation (ρ = -.17) across years between estimated GEC and the state's unemployment rate, while the correlation between the variance of the GEC estimates and the level of the state unemployment rate is positive ($\rho = .07$).[8] Nevertheless, our results suggest that the growing variance in earnings evident in unadjusted data is primarily a mixture problem, that is the growing inequality of earnings of women working full-time is attributable to growing variation in the characteristics that produce differences in expected return from work or to growing variation in the earnings payoff associated with the various factors identified in the regression.

Finally, while it is not our intention in this paper to evaluate human capital strategies for improving earnings, we note our regressions are not very encouraging on this score. In Figure 5 we plot calculated GEC for our representative woman under three alternative assumptions about education: (1) that she has completed 11 years (i.e., less than high school), (2) that she has a high school degree, and (3) that she has a college degree. We continue to assume that she is 25 years old, so the progression from high school degree to college degree reduces experience from 7 to 3 years. Clearly, while moving from less than a high school degree to diploma status leads to higher predicted earnings, and the gap between these statuses has grown, the downward trend in potential earnings is present both for those with and without a high school degree.

Net Economic Capacity

The final adjustment is to account for changes in tax liability. Median NEC was computed on the basis of applicable tax law. Appendix 2 shows each component used in our calculation of the two measures. We did not include California taxes because a woman in the circumstances we have constructed would not be liable for them. Our measures of median GEC and NEC are plotted in Figure 6.

The data in Appendix 2 and Figure 6 illustrate tax trends over the period. The upward trend in taxes in the 1970s can be seen by comparing 1972 and 1979. Although real GEC declined by over

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[8]It is possible that the positive negative correlation between GEC and state unemployment rates is attributable to cyclical variability of hours within the “full-time” category. Recall all persons working 35 hours or more per week are classed as full time.

FIGURE 5
Education and Gross Earnings Capacity

Estimated for Representative California Woman, 1972-1994

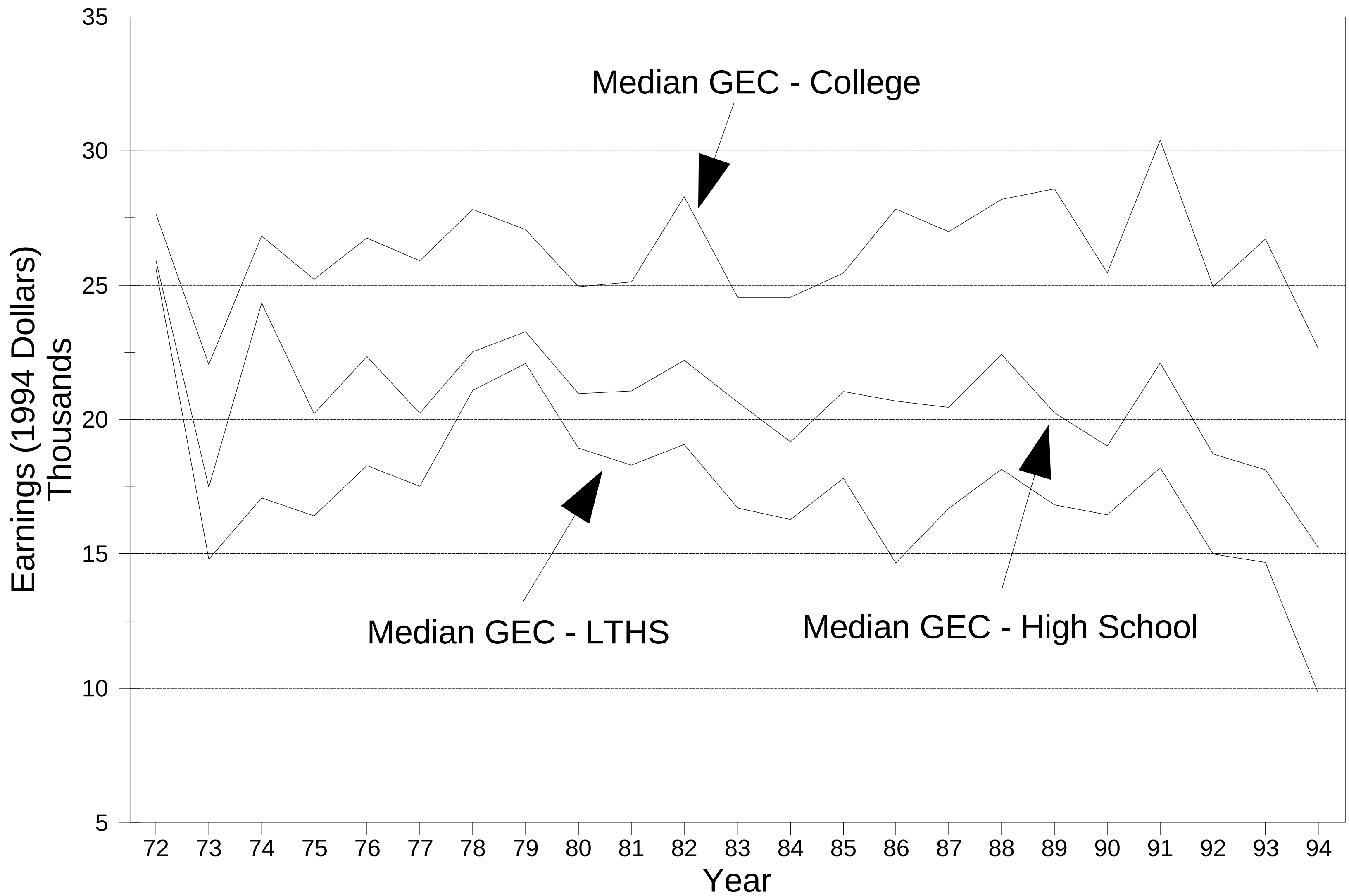
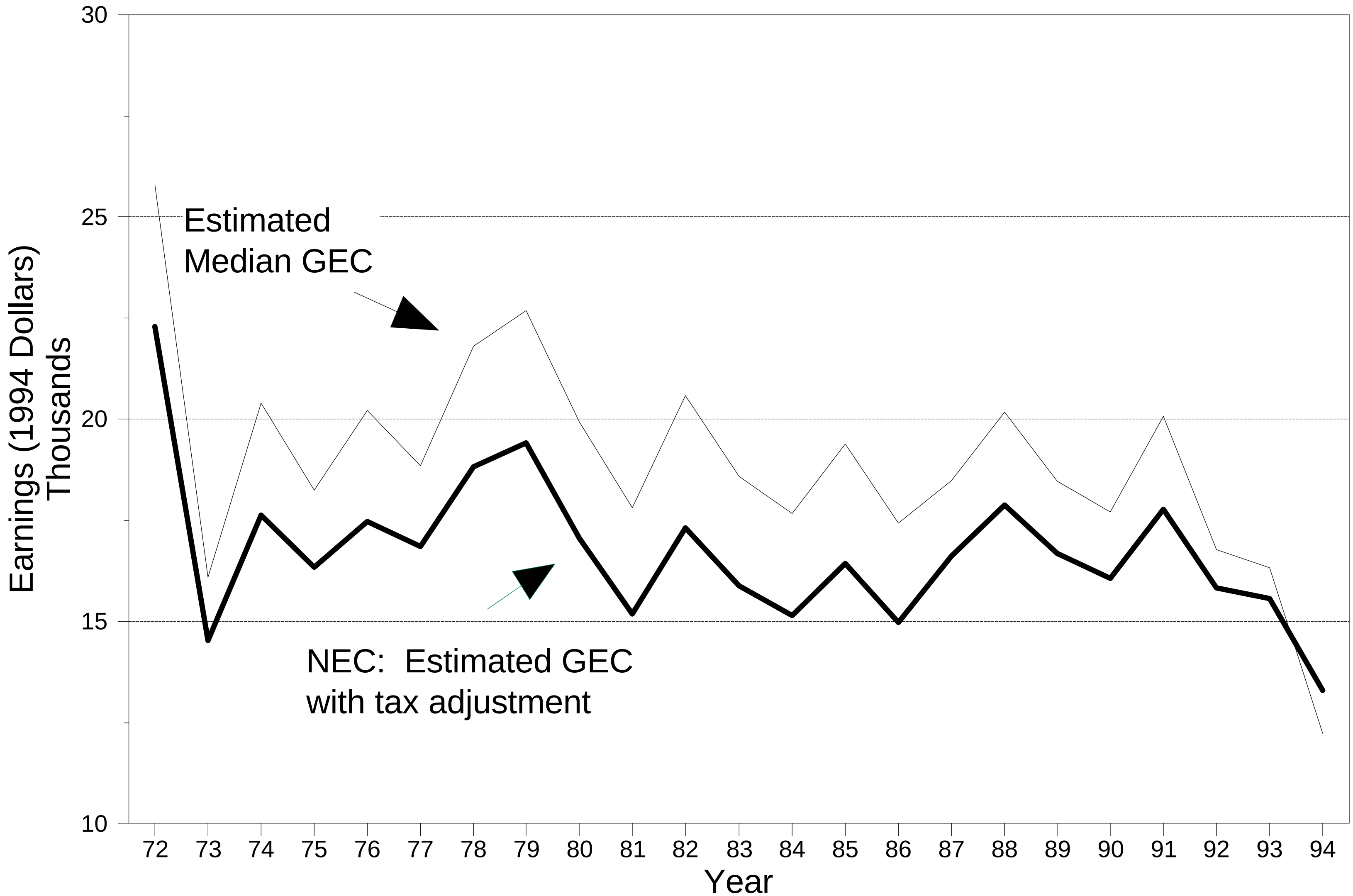


FIGURE 6
Taxes and Estimated Earnings Capacity
Representative California Woman, 1972-1994



$3,000, the average tax rate increased from 13.6 percent to 14.4 percent. After 1979, real GEC began to fall, but average tax rates remained above 14 percent until the provisions of the Tax Reform Act of 1986 became effective in 1987. After 1986, lower federal income taxes and the EITC offset the small increase in Social Security taxes, and real GEC grew until 1991.

The downward trend in average tax rates can be seen by comparing 1984 and 1990. In both years the representative individual is estimated to receive approximately the same GEC, yet the average tax rate declined from 14.3 percent to 9.2 percent. The increase in the EITC after 1992, combined with lower earnings, dramatically lowered average tax rates, to the extent that the representative woman would receive a net tax credit in 1994. A linear regression of GEC on time reveals a decline of about 22 percent (.96 percent per year) over the entire 1972–94 interval. A linear regression of NEC on time reveals a decline of about 16.4 percent (.71 percent per year) over the entire 1972–94 interval. The favorable impact of taxes on trend is largely attributable to the expansion of the Earned Income Credit contained in the Omnibus Budget Reconciliation Act of 1993, with some decline also due to the Tax Reform Act of 1986.

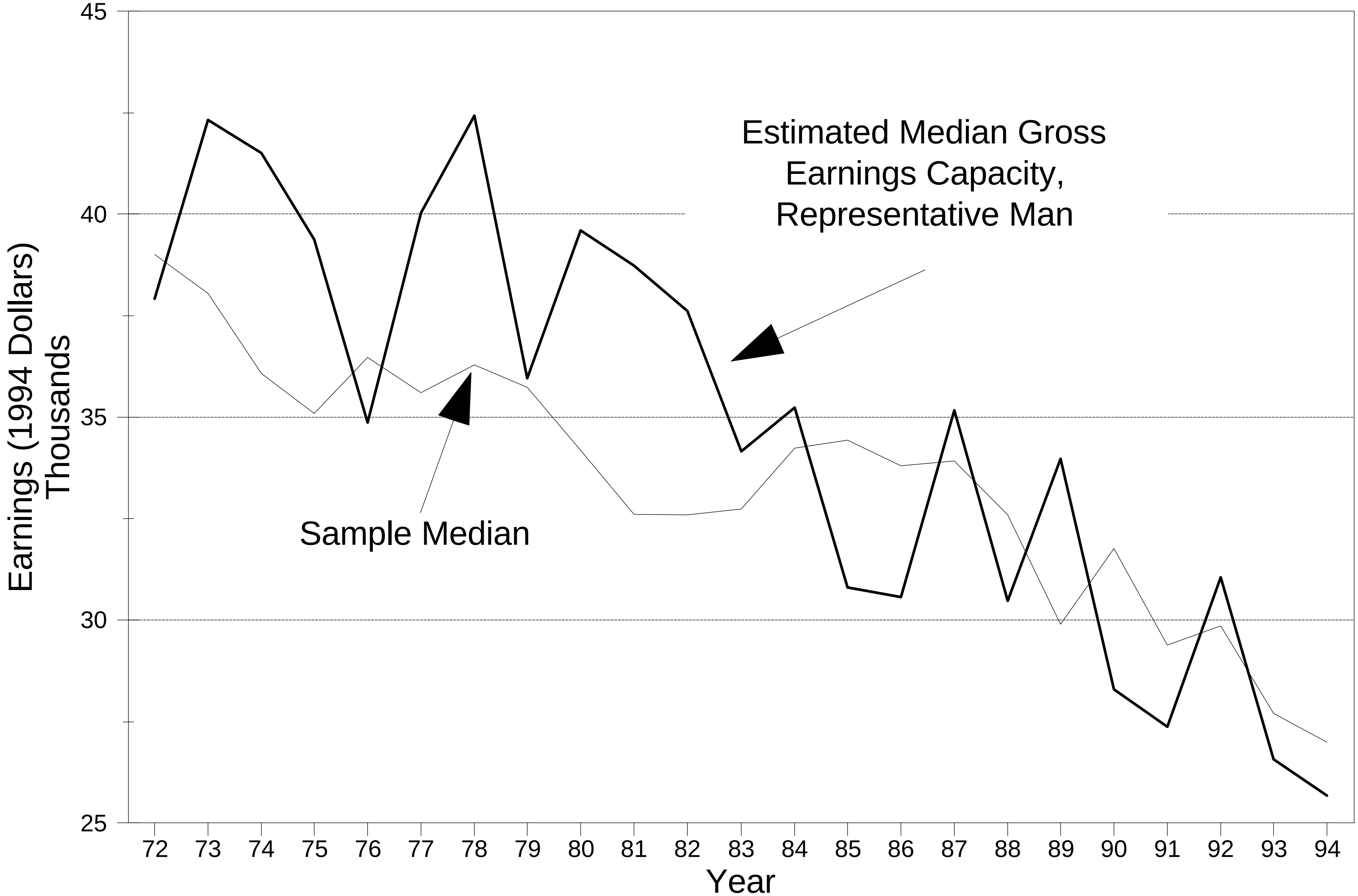
THE RESULTS FOR MEN

For purposes of comparison, we have repeated our method using men rather than women. The results for the 1995 CPS are reported in Appendix 3. Figure 7 is the equivalent for men of Figure 4. As comparison of the two figures indicates, selection and mixture adjustment has much less effect for men than was the case for women. However, the conclusion is the same: That earnings prospects for relatively low-skilled workers in California have declined substantially over the past decade. In the men's case the decline from 1988–90 to 1993–95 continued a trend established at the beginning of the 1980s. The change—wages down by 11 percent—is only slightly less than that experienced by women. The cumulative change since 1980 is substantially greater.

FIGURE 7

Estimated GEC and Median Earnings

Representative California Man Working Full-Time, 1972-1994



COMPARISON WITH OTHER WORK

Haveman and Buron (1993) use earnings capacity to relate changes in earnings to changes in poverty. They define families as “Earnings Capacity Poor” if they are unable to generate enough income to lift themselves out of poverty, even if all working-age adults in the family work full time, year-round. They find for households at risk of welfare receipt, namely female-headed households, *earnings capacity poverty* incidence rates and official poverty rates decreased between 1973 and 1988. For all female-headed households, earnings capacity poverty rates declined 26 percent, while official poverty rates decreased 9 percent. For female-headed households with children, earnings capacity poverty rates declined 19 percent, while official poverty rates remained unchanged. Female-headed households of all races except white saw a decline in earnings capacity poverty, i.e., an increase in earnings capacity. The implication is that were employment rates to rise, poverty would fall.

Using the same time frame, the results of this investigation also show a substantial increase in earnings capacity. Mean earnings capacity for the “representative individual” increased over 20 percent, while median earnings capacity increased over 25 percent. However, as Figure 7 indicates, these results are very sensitive to the years chosen. If 1972 were selected as the base period for comparison, both mean and median predicted earnings would have declined. If 1974 is selected as base year there would be little change evident. This caveat applies to all comparisons based on two years. Since the Haveman-Buron sample is national and therefore larger than the one used here, it is possible that sensitivity of results to choice of year is diminished.

Three other recent studies measure wage and earning trends over this time period, although none does so on the basis of a measure of earnings capacity. Blank (1995) looks at changes in mean weekly earnings of women between 1969 and 1989. She finds that earnings gains vary greatly by education status. Women with less than 12 years of education have nearly no movement in real earnings over the period, with earnings increasing 1.4 percent between 1969 and 1979 and declining 0.4 percent between 1979 and 1989.

Women with a high school education saw real earnings increase only 0.7 percent between 1969 and 1979, but saw a 10 percent real gain between 1979 to 1989. The largest earnings gains are for women with more than 12 years of education. They saw a decline in earnings of 1.1 percent between 1969 and 1979, but a 25.1 percent increase between 1979 and 1989.

In contrast to the Blank results, studies by Bernstein and Mishel (1993) and Burtless (1995) identify a large negative movement in wages and earnings after 1979. Bernstein and Mishel use CPS data to examine hourly wages for women with a high school education or less in several years between 1979 and 1993. They find hourly wages fell steeply between 1979 and 1989, and continued to decline thereafter. For females 16 to 25, wages fell between 16 and 17 percent between 1979 and 1993 for most education and race categories. For females 26 to 35, wage changes varied depending on education and race. Wage declines for high school graduates ranged from 4.9 percent for whites to 18.1 percent for blacks. For women with less than a high school education, wage declines ranged from 11.8 percent for whites to 18.5 percent for blacks. Burtless reports that between 1979 and 1989 average full-time earnings received by young female dropouts fell 10 percent.

The outcome of this study is closer to results of Bernstein and Mishel. For the “representative individual,” mean earnings declined 14.4 percent between 1972 and 1979 and decreased an additional 16.5 percent between 1979 and 1989. In general it appears wages have declined for women with less than a college education. Although not shown, our estimates for a woman with a college education show a decline of 4.8 percent between 1972 and 1979 and an increase of 8.3 percent between 1979 and 1989.

Our results imply that the “gender gap” between men’s and women’s hourly (and as a result, full-time, full-year) earnings has declined among workers with comparable secondary educational attainment. This is consistent with the results of analysis of national data (cf. Blau and Kahn 1997) and aggregate data on the distribution of hourly earnings in California (cf. Reed, Haber, and Mameesh 1996, Figures 3.3 and 3.7).

SUMMARY

In this paper we have developed measures of potential earnings for men and women in California over the period 1972–94. The results indicate that, other things equal, over this interval the potential earnings of women in the state have declined slightly and the potential earnings for men have declined substantially. Trends differ among groups, with earnings for those with only a high school education or less falling relative to earnings for those with post–high school education.

The results are for full-time, full-year workers. They do not necessarily reflect trends in entry wages, and they do not reveal any differential between new and continuing workers or between part- and full-time workers in access to nonwage benefits such as health care. It is our impression that the availability of such benefits has fallen over time; if this is true, their inclusion would strengthen our argument.

These trends pose problems for attaining the objectives of the new round of welfare reform. The decline in potential earnings makes it increasingly difficult to assure that working families are better off than those receiving assistance as required by California’s equity principle. The substantial decline in the potential earnings of men suggests that marriage is a less certain route out of poverty than may have been the case in the past. And there is little evidence to suggest that even when employment is obtained, time on the job quickly pays off in terms of substantial wage increases.

Recent reforms in California’s welfare law address some of these problems. While benefits have fallen sharply (see Figure 1), the state has liberalized eligibility and benefits calculation procedures to enhance the opportunity for households to combine work with AFDC benefits to assure that employment provides earnings gains. However, these policies have not incorporated time limits on income from welfare. Retaining both the principle that assistance will be transitory and that those who work will be better off than those who receive assistance may not be possible in the context of generally declining wages. At this writing the state’s planners have yet to address the problem.

APPENDIX 1
Variable Definitions in the Earnings Capacity Equations

The sampled universe is civilian California women and men aged 18 to 44. Data are from the March CPS for 1973–1995. Those self-employed or who missed work because of illness or school attendance are not included.

Variable	Definition
FTYR	Dependent variable in Probit estimation: = 1 if respondent works 35+ hours a week for 50+ weeks, else 0.
LGEARN	Dependent variable in earnings estimation: the natural logarithm of total earnings for the year.
AGE	Age of respondent in years.
AGE24	= *max*{AGE - 24, 0}
AGE34	= *max*{AGE - 34, 0}
ED	Number of years of education completed.
HSGRAD	= 1 if ED $\geq$ 12, else 0.
HSINT	= *max*{ED - 12, 0}
COLLGRAD	= 1 if ED $\geq$ 16, else 0.
COLLINT	= *max*{ED - 16, 0}
EXPER	= max{AGE - ED - 6, 0)
EXP2	= $(\text{EXPER})^2$
ASIAN =	1 if race/ethnicity reported as Asian, else 0. Not available before 1989.
HISPANIC	= 1 if race/ethnicity reported as Hispanic, else 0.
BLACK	= 1 if race/ethnicity reported as African-American and not Hispanic, else 0.
OTHER	= 1 if race/ethnicity reported not Hispanic, white, black, or Asian, else 0.
CITY	= 1 if residence reported as central city, else 0.
SUBURB	= 1 if residence reported as balance of MSA, else 0.

NOTID = 1 if, for confidentiality reasons, CPS does not identify residence location, else 0. Not available before 1986.

HEALTHPG = 1 if respondent participates in a disability program, else 0.
Program participation includes any of the following:
(1) Receives social security or railroad retirement benefits and
    (a) is not in school, is age 19–22, and is not widowed, divorced, or separated with Dependent children; or
    (b) is age 23–59, and is not widowed, divorced, or separated with Dependent children.
(2) Receives Supplemental Security Income (SSI).
(3) Receives workers' compensation.
(4) Receives veteran disability benefits, is a veteran, and is not in school.

MARRIED = 1 if respondent married and not separated, else 0.

PREVMAR = 1 if respondent currently unmarried and widowed, divorced or separated, else 0.

NUMCHILD = Number of own children under 18 years of age.

APPENDIX 2

Calculations of Before and After-Tax Earnings Capacity for the Representative Female*

FEMALE HEAD OF HOUSEHOLD WITH TWO DEPENDENT CHILDREN

Year	Nominal GEC	FEDTAX	FICA	EITC	Nominal NEC	Real (1994) GEC	Real (1994) NEC	Tax as Percent of GEC
1972	7,276	611	378	0	6,287	25,795	22,289	13.6%
1973	4,820	184	282	0	4,355	16,087	14,535	9.6%
1974	6,784	521	397	0	5,865	20,394	17,632	13.5%
1975	6,617	440	387	138	5,930	18,228	16,336	10.4%
1976	7,763	629	454	24	6,703	20,219	17,460	13.6%
1977	7,702	425	451	30	6,891	18,835	16,852	10.5%
1978	9,592	769	580	0	8,277	21,802	18,813	13.7%
1979	11,110	923	681	0	9,510	22,680	19,412	14.4%
1980	11,080	914	679	0	9,486	19,928	17,062	14.4%
1981	10,923	875	726	0	9,311	17,809	15,180	14.8%
1982	13,402	1,243	898	0	11,265	20,582	17,301	15.9%
1983	12,486	974	837	0	10,674	18,578	15,882	14.5%
1984	12,389	906	867	0	10,614	17,671	15,139	14.3%
1985	14,069	1,150	992	0	11,932	19,378	16,434	15.2%
1986	12,890	894	922	0	11,078	17,429	14,979	14.1%
1987	14,160	541	1,012	127	12,736	18,473	16,615	10.1%
1988	16,098	874	1,209	248	14,260	20,174	17,871	11.4%
1989	15,448	731	1,160	389	13,943	18,468	16,668	9.7%
1990	15,596	701	1,193	466	14,165	17,690	16,067	9.2%
1991	18,443	1,046	1,411	346	16,329	20,069	17,769	11.5%
1992	15,766	544	1,206	868	14,886	16,759	15,824	5.6%
1993	15,914	514	1,217	994	15,178	16,322	15,567	4.6%
1994	12,228	0	935	1,995	13,287	12,228	13,287	-8.7%

(table continues)

APPENDIX 2, continued

Calculations of Before and After-Tax Earnings Capacity for the Representative Male*

SOLE EARNER MARRIED MALE WITH TWO DEPENDENT CHILDREN

Year	Nominal GEC	FEDTAX	FICA	EITC	Nominal NEC	Real (1994) GEC	Real (1994) NEC	Tax as Percent of GEC
1972	10,692	1,017	556	0	9,119	37,908	32,332	14.7%
1973	12,679	1,338	742	0	10,600	42,321	35,380	16.4%
1974	13,808	1,558	808	0	11,443	41,508	34,397	17.1%
1975	14,295	1,602	836	0	11,857	39,378	32,662	17.1%
1976	13,387	1,430	783	0	11,174	34,867	29,103	16.5%
1977	16,368	1,637	958	0	13,774	40,029	33,684	15.9%
1978	18,666	2,143	1,129	0	15,394	42,428	34,991	17.5%
1979	17,613	1,764	1,080	0	14,770	35,954	30,150	16.1%
1980	22,016	2,749	1,350	0	17,918	39,597	32,226	18.6%
1981	21,536	2,634	1,432	0	17,470	35,112	28,483	18.9%
1982	24,495	3,011	1,641	0	19,843	37,618	30,474	19.0%
1983	22,950	2,411	1,538	0	19,006	34,149	28,280	17.2%
1984	24,702	2,613	1,729	0	20,365	35,234	29,048	17.6%
1985	22,364	2,095	1,577	0	18,696	30,802	25,751	16.4%
1986	22,603	2,060	1,616	0	18,926	30,564	25,591	16.3%
1987	26,956	2,066	1,927	0	22,965	35,166	29,959	14.8%
1988	24,311	1,729	1,826	0	20,759	30,467	26,015	14.6%
1989	28,417	2,284	2,134	0	24,000	33,972	28,692	15.5%
1990	24,946	1,691	1,908	0	21,343	28,295	24,208	14.4%
1991	25,149	1,624	1,924	0	21,598	27,367	23,502	14.1%
1992	29,210	2,104	2,235	0	24,874	31,051	26,441	14.8%
1993	25,902	1,549	1,981	0	22,375	26,565	22,948	13.6%
1994	25,667	1,429	1,964	0	22,276	25,667	22,276	13.2%

Source: Authors' calculation using the following sources: 1. **FEDTAX**: computed using yearly tax tables from each year's IRS Form 1040. 2. **FICA**: Committee on Ways and Means 1994, p. 76. 3. **EITC**: Committee on Ways and Means 1994, p. 700.

*See Table 3.

APPENDIX 3

Male Probit Equation Estimation Results
Probability of Full-Time Full-Year Employment, 1995

Variable	Mean	Std. Dev	Coefficient	Std. Error
CITY	0.3646	0.4814	0.3666**	0.1630
SUBURB	0.5741	0.4946	0.5102***	0.1607
NOTID	0.0287	0.1669	-0.1599	0.2272
BLACK	0.0467	0.2111	-0.4909***	0.1282
ASIAN	0.0935	0.2913	-0.0100	0.0893
HISPANIC	0.3646	0.4815	0.0814	0.0626
OTHER	0.0148	0.2913	0.1307	0.1998
ED	12.5644	3.4336	0.0332**	0.0133
HSGRAD	0.7986	0.4011	0.0750	0.0953
COLLGRAD	0.2564	0.4368	0.2572***	0.0856
AGE	32.9938	6.4907	0.1435***	0.0117
AGE24	9.2073	6.1242	-0.1086***	0.0136
AGE34	2.2387	3.2345	-0.0477***	0.0090
MARRIED	0.5543	0.4972	0.3299***	0.0323
PREVMAR	0.0956	0.2942	0.1262**	0.0568
NUMCHILD	1.0206	1.2363	-0.0332***	0.0115
HEALTHPG	0.0051	0.0711	-0.5037***	0.1288
CONSTANT	1.0000	0.0000	-4.1213***	0.3413

Dependent Variable: FTYR
Number of Observations: 2166
Mean of Dependent Variable: 0.6758
Standard Deviation of Dependent Variable: 0.4682

*** Significant at the 1% level; ** significant at the 5% level; * significant at the 10% level.

(table continues)

APPENDIX 3, continued

Male Earnings Regressions Results
Earnings from Full-Time Full-Year Work, 1995
(Heckman Selection Model)

Variable	Mean	Std. Dev	Coefficient	Std. Error
CITY	0.3541	0.4784	0.0885	0.0874
SUBURB	0.6008	0.4899	0.1332	0.0862
NOTID	0.2037	0.1553	0.3711***	0.1219
BLACK	0.0467	0.2111	0.0389	0.0687
ASIAN	0.0935	0.2913	-0.1515***	0.0458
HISPANIC	0.3646	0.4815	-0.2664***	0.0326
OTHER	0.0148	0.1206	-0.0869	0.1066
ED	12.5644	3.4336	0.0242***	0.0078
HSGRAD	0.7986	0.4011	0.1218**	0.0503
HSINT	1.4895	1.8353	0.1412***	0.0137
COLLGRAD	0.2564	0.4368	-0.2984***	0.0605
COLLINT	0.1030	0.3803	0.0032	0.0384
EXPER	14.4294	6.9296	0.0430***	0.0038
EXP2	256.1958	219.3413	-0.0007***	0.0001
CONSTANT	1.0000	0.0000	9.5131***	0.1210
LAMBDA	—	—	-0.5465***	0.0655

Dependent Variable: LGEARN
Number of Observations: 1491
Mean of Dependent Variable: 10.1839
Std. Deviation of Dependent Variable: 0.7163
Log Likelihood: -3324.50
R^2 (unadjusted): .41

*** Significant at the 1% level; ** significant at the 5% level; * significant at the 10% level.

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